Father as Breadwinner:

Gendered Wage Penalties for Job Interruptions*

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Abstract

The wage penalty associated with motherhood and women's greater job discontinuity is well-established. However, wage outcomes associated with men's family status and job discontinuity is less well understood. This study examines the effect of job interruptions on wages for both men and women and disaggregates job interruptions into reasons for them in order to examine specific wage effects associated with different types of leave. Further, the possibility that men incur wage penalties associated with parenthood in the form of penalties associated with job absence is considered. This study examines both immediate and longer-term effects of job interruptions on women's and men's wages over time. Using a partial-adjustment model based on fixed effects estimates, data from the NLSY(1979-1998) are used to estimate changes in wage trajectories as a function of family status and job interruptions. Results show that women experience greater job interruptions due to family reasons compared to men, and also receive wage penalties for such interruptions net of human capital and the direct effect of children. Men appear to get a "pass" for family job absences, as no wage penalties associated with such absences are found for men. Penalties to women's wages associated with voluntary and family-related job interruptions are exacted in the short-term, but effects do not persist over the long run. Compared to women, men incur larger wage penalties associated with non-family job absence, and these penalties do remain over the long-run. Women's wages are depressed directly by the presence of children, while men's wages are penalized by job interruptions. Results implicate the reinforcement of men's "breadwinner" work and family role as a primary obstacle to gender equity in the labor market.

Introduction

The "price of motherhood" has recently received considerable academic and popular attention (e.g., Budig and England 2001; Anderson, Binder, and Krause 2003; Crittenden 2001). However, the "price of fatherhood" has been given far less consideration. Given the decline in traditional gender ideologies over the past half-century (Ciabattari 2001; Thornton and Young-DeMarco 2000), women's increased labor force participation (Blau 1998; Cohen and Bianchi 1999; Goldin, 1990), and a decline in the number of gender-specialized households composed of a full-time working man and unemployed woman (Bianchi forthcoming; U.S. Census Bureau 2000), the relative dearth of research on men's work-family issues warrants rectification. Thus, this study examines the "price of parenthood" and includes men in the work-family research conversation.

Despite recent attitude shifts away from gender differentiation, gender inequality in the labor market persists. Beyond the well-established fact that women earn less than men[1], marriage significantly decreases women's wages (Chandler, Kamo and Werbel 1994), increases women's job interruptions (Gerstel and McGonagle 1999), and limits women's job mobility (Felmlee 1984; 1995). There are no such penalties for men. To the contrary, married men earn substantially more than unmarried men and this pay premium accounts for approximately one-third of the estimated gender wage discrimination (Bartlett and Callahan 1984; Gorman 1999; Korenman and Neumark 1991; Lundberg and Rose 2002; Neumark 1988).

Particularly acute is the effect of children on women's wages. Referred to as the "motherhood penalty", the research literature consistently reveals a wage penalty of between 4 to 15 percent for women with children, as compared to women without children (Anderson, et. al 2003; Budig and England, 2001; Fuchs 1988; Korenman and Neumark 1992; Peterson 1989;

Waldfogel 1997), and this penalty has persisted over the past two decades (Avellar and Smock 2003; Waldfogel 1998a). There is no “baby tax” for men. To the contrary, some studies find a positive effect of children on men’s wages (Lundberg and Rose 2000; Waldfogel 1998a).

Moreover, the presence of children increases absenteeism for women, but not for men (Leigh 1983). Women with children have higher incidents of job interruptions than men and women without children, and women’s conflicting work and family responsibilities continues to be an important element contributing to gender earnings inequalities (Albrecht, Edin, Sundstrom and Vroman 1999; Goldin 1990; Jacobsen and Levin 1995; Waldfogel 1998a). Men, on the other hand, respond to the birth of a child in quite an opposite way, and are more likely to *increase* their labor supply, rather than take time off from work (Lundberg and Rose 2000; Mennino and Brayfield 2002). Qualitative research evidence suggests that because men’s work roles are central to their family roles as breadwinners (Gerson 1993; Pleck 1993), an appropriate male response to family demands such as the birth of a child, is to increase their work hours even further, in simultaneous fulfillment of their overlapping work and family roles (Kaufman and Uhlenberg 2000; Mennino and Brayfield 2002). Qualitative and anecdotal evidence further implicate adherence to gender norms in men’s reluctance to take leaves from work for family reasons, as a pervasive perception of employer disapproval of men’s leave-taking prevents men from doing so (Pleck 1993; Williams and Segal 2003). Thus, if men’s primary work and family roles both define him as “breadwinner”, then an absence from work represents a transgression of these normative roles. As a result, men may incur a “parenthood penalty” through penalties associated with job interruptions, rather than through a direct effect of children.

A large body of research documents the negative effects of job interruptions on women’s earnings (Albrecht, et. al 1999; Felmlee 1984; Jacobsen and Levin 1995; Lundberg and Rose

2000), labor supply (Drobnic, Blossfeld, and Rohwer 1999; Lundberg and Rose 2000), job mobility (Felmlee 1995), and career trajectories (Rosenfeld and Spenner 1988). However, studies examining effects of job interruptions on labor market outcomes focus almost exclusively on women, and these analyses focus almost exclusively on family factors even if specific reasons for job absence are not made explicit (Rosenfeld 1992). This study unpacks job interruptions by identifying reasons behind them, for both men and women. In doing so, effects of job absence on wages may be more precisely described, and the possibility that a "parenthood penalty" exists for men in the form of wage penalties associated with family leave is examined.

Given trends in attitude shifts toward gender egalitarianism in work and family, coupled with recent innovations in family policy that provide for gender-neutral parental leave (the Family and Medical Leave Act of 1993), research on men's work-family negotiations and outcomes as well as women's is timely. Drawing on data from the National Longitudinal Survey of Youth (1979-1998), this study estimates a partial adjustment model of change to address two primary questions to further our understanding of persistent gender inequality in the labor market: (1) Do men incur the same long-term wage penalties associated with family status (marital status and the number of children) and job interruptions as women do? And (2), net of family status and labor force experience, is the penalty for job interruptions distributed equally between types of absence from work, or does a "parenthood penalty" manifest in a larger wage penalty specifically associated with taking a family leave relative to other types of job absences? The partial adjustment model allows an estimation of long-term effects of family status and job interruptions on men's and women's highest expected potential wage, which are then used to characterize wage trajectories over time.

Background and Significance

Although women's increased participation in the labor market over the past century has not been matched by a commensurate increase in men's participation in household labor, men have been gradually increasing their share of household and family labor over the past several decades. Recent research finds a slight but significant increase in married men's contribution to housework (Bianchi, Milkie, Sayer and Robinson 2000; Brines, 2000; Casper and Bianchi 2001), and a significant increase in fathers' involvement in childcare since the late 1960s (Bianchi 2000; Sandberg and Hofferth 2001). Fathers are also reporting a desire to spend more time with their children (Bianchi forthcoming; Gerson 1993; Hyde, et al. 1993), and the majority of both men and women today want fathers to be equally involved in child care (Milkie, Nomaguchi and Bianchi 2002). Although a gap between reported intentions and behavior remains, some evidence does also suggest a slight increase in men's interruptions from work in the form of family leave-taking (Bond, et al. 1991; Commission on Leave 1996; Han and Waldfogel 2003).

While considerable research efforts have been devoted to understanding wage consequences of women's job interruptions associated with family reasons, much less is understood regarding consequences to men's wages. Available research on housework indicates that men's participation in housework results in negative labor market outcomes similar to women's experience (Noonan 2001; Coltrane 1996; Coverman 1983; Gerson 1993). Limited research on job interruptions suggests the same. Data from the 1973 Quality of Employment Survey suggest that men's and women's absence rates respond similarly to endowments such as wages and responses to endowments such as child care (Leigh 1983). Another study of Australian employees, however, indicates structural differences in men's and women's absence behavior, such as the commute to work, shift work, union membership, occupation, and industry

(VandenHeuvel and Wooden 1995). Focusing only on job absences due to illness or injury, Vistnes (1997) finds that among men who have missed work, those with young children in day care miss 40 percent more days from work than men whose children are not in child care. The author suggests that men with children in day care are likely to have working wives, and that these fathers behave more like working mothers than like other men.

The positive effect of a continuous work history on earnings has received empirical support in several studies. Prior research suggests that a continuous work history (years out of work, years with current employer, etc) and indicators of work force attachment (limits on job hours or location, hours of work lost due to job absence) explain 42 percent of the wage gap between white men and white women (Corcoran and Duncan 1979). More recent studies find that even a single break in employment generates immediate costs to women's wages and occupational attainment (Jacobsen and Levin 1995), and women with gaps in their employment earn lower wages upon reentry than at the time of withdrawal from the labor market (Klerman and Leibowitz 1999). Moreover, such penalties associated with women's job interruptions are cumulative, resulting in long-term consequences of depressed wage growth over the working life course (Felmlee 1995; Mincer and Ofek 1982). Parallel effects of such job interruptions on men's wages remain less well understood.

Women's greater likelihood of career interruptions relative to men's is implicitly understood as a consequence of women's unequal share of family responsibilities, and such interruptions are considered a major contribution to the gender gap in wages (Jacobsen and Levin 1995; O'Neill and Polachek 1993). A body of work relying largely on Becker's (1981) model of household decision-making posits that within a marital household, the responsibility for wage-earning will be primarily allocated to the marital partner earning higher wages. Because the husband usually out-earns the wife, he will bear primary responsibility for wage earning, while the responsibility for household production is absorbed by the wife. As a result, women are more likely to interrupt their employment to care for family and household responsibilities.

If human capital accumulates with work experience (Becker 1981; O'Neill and Polachek 1993), and if women experience more interruptions in their work history than men, they will accumulate less human capital, and their earnings profiles are likely to be flatter as a consequence. According to human capital theory, women's intermittent labor force participation, as well as their lower investment incentives in anticipation of intermittent employment due to family responsibilities, leads to lower earnings and a flatter earnings profile relative to men (Becker 1981; Fuchs, 1989; Polachek, 1995). This paper examines the effects of job interruptions on men's wages as well, in order to determine whether wage effects associated with job interruptions are gendered. That is, do men and women receive similar wage penalties to job interruptions, or do women receive greater wage penalties due to employer bias or lowered productivity? The idea that women's lower wages as a result of family-related job interruptions is tested in this paper, and in addition, this paper asks whether men too, experience a parenthood penalty in the form of penalties specifically associated with job interruptions for family reasons.

Previous studies examining consequences of job interruptions on labor market outcomes have not differentiated between job absences associated with family responsibilities (such as pregnancy and child care needs) and other types of absence from work. The few studies that do differentiate between types of job absence focus instead on the interruptions themselves, and these studies indicate that the majority of women's job exits are *not* in fact related to marital or family status. Using data from the NLS Young Women, Glass (1988) reports that while 13 percent of job exits are associated with marriage and 20 percent with pregnancy, 67.5 percent of women’s job exits are not associated with family transitions. Similarly, job exits due to a very loose definition of “family reasons” amount to only 39 percent of women’s job exits in Wenk and Rosenfeld’s (1992) analysis of women in Washington State. Using national data, Vistnes (1997) finds no relationship between women’s likelihood of absence and child care needs. And, Taylor (1981) reports that both men and women lose more days of work due to illness than for any other reason. These studies suggest that while women may experience more job interruptions than men due to their role in the family, most of women’s job absences are for reasons similar to absences experienced by men. The consequences of these different types of absences, however, have not been teased apart.

This study unpacks the effect of job interruptions on wages and identifies three different categories of job interruptions, according to reasons given for the absence from work: family leave, non-family absence, and involuntary absences from work. Such categorization of interruptions provides a more complete and accurate explanation of the wage process and penalties associated with specific job interruptions. If, as human capital theory suggests, women with children earn lower wages as a consequence of increased job absences that lead to skill depreciation, the effect of job interruptions on wages would explain the parenthood penalty associated with the presence of children. Further, wage penalties associated with job absences should be equally distributed between types of leave; it should not matter what caused the absence from work. And, theoretically, men’s and women’s wages would be expected to respond in the same way as each other in response to job interruptions. Wage penalties distributed differently across different types of interruption or differently between men and women would suggest that some other mechanism may be operating to drive wage penalties.

Studies examining the “motherhood penalty” generally attribute residual unexplained association between parenthood and wages to some combination of employer discrimination or differences in parent versus non-parent productivity (Budig and England, 2001; Lundberg and Rose 2000; Waldfogel 1997b). Although not directly measurable, labor market discrimination is estimated to have a substantial negative effect on women’s earnings (Blau and Ferber 1987). If employers value and reward workers who are perceived to be consistent employees and likely to remain with the firm (England and Farkas 1986), married women and women with children may be presumed to be at greater risk of leaving the firm or taking time off to tend to family responsibilities, and are therefore, less valued (Peterson 1989; Siebert and Sloane 1981; Waldfogel 1998a). By the same token, employers may place a higher value on married men with children who may be perceived to be better able to provide stable employment as the primary financial providers for their families (Hersch and Stratton 2000). Further, these men may be presumed to also have wives at home to absorb the costs to time on the job that are associated with tending to family responsibilities (Becker 1981; Korenman and Neumark 1991). These gendered employment practices put women at a disadvantage and lead to earnings inequalities.

In addition to gender differences in employer biases, different types of absence from work may also be associated with different employer evaluations. While some types of absence from work are understood as entitlements, such as vacation and sick leave, family leave-taking may be seen as less acceptable to employers, despite policies that provide for it. Employees who take family leave may be inadvertently ‘signaling’ weaker commitment to their job, and this may suggest to an employer a lower capacity for productivity or the potential for future absence from work (Albrecht, et al 1999; Jacobsen and Levin 1995). Moreover, different types of leave may carry different consequences for men and women. If employers indeed prefer male workers premised on the assumption that men are less likely to have job interruptions, then this normative expectation of men's primary breadwinning role may lead to greater wage penalties associated with family leave-taking for men, as taking time off from work to care for family needs runs counter to normative expectations of male work and family roles, and may be discouraged by employers. A qualitative study of 138 men implicates social disapproval and structural barriers at work for men's unequal contribution to family care (Gerson 1993). In another study of 550 husbands of pregnant women in four Wisconsin health clinics, 43 percent anticipated negative co-worker reactions to the respondent's family leave-taking (Hyde, et al. 1993). Sixty-three percent believed their supervisors would react negatively to their taking family leave. Even in Sweden, where parental leave policies are most generous and 100% compensated, men's perception of employer disapproval for transgressions of normative gendered expectations is consistently cited for their low participation rates in parental leave programs (Albrecht, et al 1999; Haas and Hwang 1995). Indeed, adherence to gender-specific role expectations in the work-family context has been found to enhance psychological well-being (Carr 2002).

The relationship between parenthood and wages has also been attributed to differences in productivity allocations between mothers and women without children (Budig and England 2001), as well as between mothers and fathers (Lundberg and Rose 2000). Women with children may conserve their energy on the job to spend with their children at the end of the day, or they may be so exhausted from working the double shift at home and work that their productivity suffers and their wages fall as a consequence. For men, as the primary breadwinners for their families, married men and men with children may be more motivated to work harder and allocate more time and effort to work, and are thereby rewarded with higher wages relative to unmarried men who may not be similarly motivated, and may not be as productive. While this study employs fixed effects to account for stable unobserved characteristics, unmeasured characteristics that change in an individual over time cannot be accounted for in the models to be tested in this study. For example, if a worker's ambition, motivation, and energy changes over time, perhaps in response to changes in family status, the models to be estimated in this study will not hold these constant.

Current Investigation

The dynamic relationship between work and family exacts wage penalties that are unequally distributed between men and women (CITE), the consequences of which are long-term and cumulate over the life course, such that women's earnings relative to men's declines with age (Marini 1989). This study examines trajectories of men's and women's wages as a function of family and marital status, job interruptions, and human capital. In projecting wage trajectories, I expect to find that the gap between men's and women's wages widens over the course of the work career. If men have more consistent work histories and fewer job interruptions, they will enter into jobs with greater opportunities for advancement. As a result, their career ladders will likely be longer than women's. Over time, men's earnings trajectories will be higher and steeper than women's, as illustrated in Figure 1.

Figure 1 about here

If, as human capital theory suggests, women earn lower wages as a consequence of increased job interruptions that result in skill depreciation, the effect of leave-taking on wages would completely explain the penalty associated with parenthood, and no independent effect of the presence of children would be found. Further, penalties associated with job interruptions should be equally distributed across types of job leaves taken, in the absence of employer preferences. Different wage effects associated with different types of leave would suggest some other mechanism, beyond the generic absence from work that drives wage penalties, and may point to alternative explanations. The possibility of a wage penalty specifically associated with family leave-taking is explored. Women and men in particular, who take such leaves may inadvertently signal conflicts inn their commitment to work and this may affect long-term wage trajectories over time for workers who have such absences relative to those who do not, despite work-family policies designed to prevent negative wage and career outcomes.

Similarly, if family status exerts an independent effect on wages net of leave-taking, the specter of employer bias would be raised, although the possibility would remain that workers who are also parents conserve their energy at work to prepare for their “second shift” at home. This could lead to lower productivity and could also explain a negative relationship between the presence of children and wages.

As traditional gender roles continue to erode and men continue to increase participation in household and family labor, more research is necessary to understand how men’s participation in traditionally women’s roles affects their market outcomes. Such research may also identify obstacles to men’s full participation in family and household labor, thereby also identifying obstacles to gender equity in labor market outcomes. This paper examines wage effects of parenthood for both men and women and estimates parameters that characterize a wage trajectory over time as a function of family status, job interruptions, and human capital.

DATA

This study uses the 1979-1998 waves of the National Longitudinal Survey of Youth (NLSY 1979-1998). The NLSY is comprised of a nationally representative sample of 12,686 young men and women aged 14-21 when first surveyed in 1979. The data were collected annually from 1979 through 1994. After 1994, the data were collected biennially. Individuals who stopped responding to the survey by the second year of the survey were excluded from my sample, as were the NLSY military and poor white supplemental samples, which were discontinued in 1984 and 1990, respectively. These restrictions reduced the sample to 9,650 respondents.

The surveys contain a core set of labor experience questions, in addition to questions regarding marital and fertility histories. Using the detailed work histories, individuals are identified as "at risk" for experiencing a job interruption beginning from the point at which individuals enter the labor force. Labor force entry is defined in this study by the year in which the respondent reports earning a wage working a minimum of 20 hours a week, while not concurrently enrolled in school. This definition is intended to begin observing individuals at the point in which their primary attachment is to the labor market, rather than to schooling.

The sample is further limited to individuals with at least two years of labor force experience between 1997 and 1998, as fixed effects models require two points of observation. This left the final sample with 9,204 cases. In the analyses that follow, yearly observations from these cases are pooled in a cross-section time series. Person-year observations with a reported hourly wage of less than \$1.00 or greater than \$400.00 were omitted. This resulted in a deletion of only 0.5 percent. Person-year observations in which a respondent was in active military duty, in jail, or was working on a farm were also excluded, yielding a total of 94,321 person-year observations. Restricting the sample in these ways likely captures individuals with stronger attachments to the labor market relative to the general population. Thus, effects of job interruptions in particular are likely to be underestimated, as persons with strong attachments to the labor market are less likely to experience job interruptions, and are less significantly affected by them (Wenk and Rosenfeld 1992).

Measures

Table 1 contains the means and standard deviations for variables used in the models, as observed in the first (1979) and last (1998) years of the survey, spanning a total of 20 years. The dependent variable is the natural log of hourly wages in the respondent's current job, converted into constant 1998 dollars.

Table 1 about here

Each of the independent variables included in the models are measured in the year prior (t-1) to the observed wage (t). This one-year lag allows some time for the measures in the model to affect wage outcomes, particularly with respect to effects associated with job interruptions.

The central independent variables of interest are those regarding interruptions in work history over an individual's working life. Retrospective survey questions in the NLSY ask both men and women whether they experienced breaks in employment either within or between jobs over each year, the duration of that break, and the reason for it. Job interruptions are measured in this study as the cumulative number of weeks a respondent reported a gap in their work history in each year. A "family job absence" is defined in this study as any break in employment due to pregnancy, child care problems, or family reasons. "Voluntary, non-family job absence" is defined as a break because the respondent felt ill or unable to work, was on vacation, was in school, did not feel like working, or "other". An "involuntary leave" is defined as an absence from work due to disability, strike, layoff, or the unavailability of work.

The means and standard deviations in Table 1 indicate a great deal of variation in leave-taking. As expected, men experience far fewer weeks of job interruptions due to family reasons than women. In the first year of labor force participation, men have an average of less than one week of family-related absence from work, compared to 5 weeks for women. In their last year of work, both men's and women's absences due to family reasons seem to decrease, and men take about half of one week off from work for family reasons, while women take about 3 weeks in their last year of employment observed. Women also experience slightly greater absences from work due to voluntary, non-family reasons, although men have a slightly higher incidence of involuntary job interruptions. One should note, however, that the standard deviations indicate substantial variation across all types job interruptions experienced.

Due to the skewed distribution of the leave variables, each has been recoded into a set of dummy variables. According to the Family and Medical Leave Act of 1993, employees are guaranteed 12 weeks of job-protected leave each year for family care. Thus, dummy variables are created out of the job interruption measures by identifying whether respondents had more or less than 12 weeks of absence from work in each year, as persons who have less than 12 weeks of absence at least for family reasons should be covered under the FMLA. The reference group for all types of job interruptions is having no absence.

Marital status is measured as a dichotomous variable (never married or ever married). Approximately one-quarter of the sample is married in 1979 and 74% are married in 1998. The number of children a respondent has is counted by the respondents' report of each child in the household roster. Thus, the children reported may be biological, and may also include adopted children or other children in the respondent's household, and presumably under the respondent's care. By 1998, the respondents in this sample have had an average of one and one-half children.

Human capital control variables include the highest year of schooling completed, weeks tenure at the current job, and labor force experience. The latter is measured separately as a running count each, of the number of weeks a respondent has reported earning a wage working full time (35 or more hours a week) and part time (less than 35 hours a week). A dummy variable indicating part-time or full-time employment of each year's current job is included, as is a dummy variable indicating whether the individual belonged to a labor union. A control labor union membership is included as previous studies have found higher rates of absenteeism among union members (Leigh 1986; VandenHeuvel and Wooden 1995). Avenues by which union membership might affect job absence include enhanced job security and bargaining power for benefits that may affect both wages and the likelihood of absence from work.

A dummy variable identifying respondents employed in professional and managerial occupations is included in each analysis, with all other occupations grouped as the reference category. Preliminary models that included dummy variables for each occupational category did not provide any additional information. Thus, the occupational categories were collapsed in this manner to produce a more parsimonious model.

In order to control for the possibility that chronic health conditions affect individuals' ability to work, and, consequently, affect their wage outcomes, a dummy variable is included in all models to identify respondents with chronic health conditions.

METHOD

First, the NLSY data are arranged into a pooled cross-section time-series in which the unit of analysis is the person-year[2]. Because previous research has found significant bias associated with selection effects (Anderson, et. al 2003; Budig and England 2001), this study estimates fixed-effects regression models to analyze the pooled cross-section time series data to control for selection. The models will be estimated in a stepwise fashion in order to more closely examine the effects of job interruptions and human capital on the relationship between parenthood and wages. In the second part of the analyses, the fixed effects coefficients are applied to a partial adjustment model of change to estimate a maximum potential wage, as well as the speed at which that wage is achieved.

The fixed-effects model is expressed as:

$$Y_{it} - \overline{Y_i} = b_0 + b_i(X_{it} - \overline{X_i}) + (\varepsilon_{it} - mean(\varepsilon_i)) \text{ , t = 1,2,….,T} \tag{1}$$

where b_i denote regression coefficients, b_0 denotes the intercept, *k* indexes the measured independent variables (*x*), *i* represents individuals, and *t* indexes time. The error term is represented by ε, the structure of which is explained in further detail below, in phase 2.

$$Y_{it} = b_0 + b_i X_{it} + \varepsilon_{it} \text{ , t = 1,2,….,T,} \tag{2}$$

In equation (2), the person-specific unobserved effects have been removed by subtracting from each variable its person mean across years, as demonstrated in equation (1). This removes potential biases in the coefficient estimates associated with fixed individual characteristics that would lead to selection bias.

In the second part of the analysis, estimates from fixed-effects regression models will be used to estimate a partial adjustment differential equation. The partial adjustment model will characterize wage trajectories that conform to the shape illustrated in Figure 1, by estimating an ultimate wage at “equilibrium” over the working life-course for individuals. This equilibrium wage represents the highest wage an individual is expected to earn as a function of the specified model parameters, at which point, subsequent earnings plateau. The model also obtains estimates of the speed at which that potential wage is reached. This estimate is not meaningfully interpretable in its own right; rather, it serves as a point of comparison between men’s and women's wage trajectories. The wage trajectories and equilibrium wages are each estimated as a function of a vector of exogenous variables representing family status and employment history.

The Partial Adjustment Model of Change

The partial-adjustment model is one of a class of differential equation models that derives parameters that characterize the trajectory of hourly wages as a function of individual characteristics. The model also generates estimates of the speed at which Y achieves its maximum potential or "equilibrium value" (Neilsen and Rosenfeld 1981). The model is formally expressed by the equation below:

$$dY/dt = -b(Y^*-Y) \qquad (3)$$

Define Y as a person's hourly wages, and Y* as the value of one's wage at the peak of his/her career[3]. In this model, the value of hourly wages during the first year of employment is assumed to be less than Y*, and equation (3) implies that the rate of change in wages (dY/dt) is proportional to the distance that remains between wages observed at the current time, and wages at equilibrium. The coefficient b characterizes the trajectory of wages over time and is constrained to be negative to represent a wage trajectory that approaches Y* at a decelerating rate. The larger the b in absolute value, the more quickly a person's wage reaches equilibrium (Y*). Y* can be defined as a linear function of a vector of exogenous variables affecting wages:

$$Y^* = a_0 + a_1x_1 + a_2x_2 + \ldots. a_kx_k \qquad (4)$$

Substituting the right side of Equation (4) for the Y* in Equation (3), and solving the differential equation for Y_t (Coleman 1968; Rosenfeld 1980), returns the following expression:

$$Y_t = e^{bt} Y_0 + a_0(1- e^{bt}) + a_1(1- e^{bt})X_1 + \ldots. + a_k(1- e^{bt})X_k \qquad (5)$$

Figure 2 depicts the approach of a predicted wage trajectory to its equilibrium point, derived from dY/dt = bY + aX. This illustrates a typical trajectory of dynamic systems characterized by b (Nielsen and Rosenfeld 1981) and corresponds to empirical observations of wage profiles. Individuals typically enter the labor market earning lower wages than those they end up with, advancing through promotions or higher-paying jobs until they reach their highest earning potential, at which point, wages plateau (Nielsen and Rosenfeld 1981).

Figure 2 about here

Next, Equation (5) is fitted with data. This is done by estimating a linear difference equation using lagged values of the dependent variable and the exogenous variables (Coleman 1968; Nielsen and Rosenfeld 1981). A one-year lag is represented in the following equation:

$$Y_t = b^* Y_{t-1} + a_{0t}^* + a_1^* X1_{t-1} + a_{1\Delta}^* X_{1\Delta} + \ldots + a_k^* X_k + e \qquad (6)$$

The estimated ("*") coefficients can be transformed to solve for the parameters of the dynamic model – that is, for b in Equation (3) and the a_i's in Equation (4). In comparing equations, b* in Equation (6) corresponds with e^{bt} in Equation (5). Thus, the estimate for b = ln(b*) and $a_1 = a_1^*/(1-b^*)$. These calculations provide estimates for the long-run effects of exogenous variables on equilibrium wages (Y*). The transformed estimates, unlike the coefficients of a directly-estimated differential equation, can be meaningfully compared across equations as coefficients of the path of wages over the working life-course, and they form the basis of comparisons by sex below.

Equation (6) may be estimated using ordinary least squares regression. However, the inclusion of lagged dependent variables may lead to violations of OLS assumptions. If there are omitted variables and those unmeasured characteristics are relatively stable or systematically related over time, the error terms will include the effects of these excluded variables, which results in biased estimates due to autocorrelation (Rosenfeld 1980).

Applying a fixed effects model reduces such bias, as it allows for correlation between the error term and explanatory variables in any time period. The structure of errors is modeled as:

$$\varepsilon_{it}, = u_i + v_t + w_{it} \tag{7}$$

where the components of the error term represented by *u* capture the cross-sectional person-specific component of error that remains relatively constant over time; *v* represents the unobserved variables that vary over time, but in the same manner across individuals; and *w* represents the usual, random error component. The resulting coefficients are those that would be obtained if dummy variables were included to represent each respondent in the sample for each year they are in the sample. Coefficients from fixed effects estimates will be applied to the partial adjustment model to estimate the effects of a vector of individual family- and job-related characteristics on wage trajectories over time.

RESULTS:

Table 2 contains the estimates from the fixed effects model on the pooled cross-section for women, and Table 3 contains estimates for men. A prior OLS regression analysis that included an interaction term for sex * wage indicated statistically significant differences in the wage process between men and women. Thus, Tables 2 and 3 contain the fixed effects estimates separately for men and women.

Table 2 about here

Contrary to some studies that find a negative relationship between marriage and wages for women, results reported in Table 2 are consistent with those reported by Budig and England (2001) and indicate that women receive approximately a 3% wage premium for marriage. Consistent with previous research, Model 1 of Table 2 also indicates an approximate 4%

decrease in women's wages for each child in the household. In Model 2, the set of job interruption measures is included and reduces the negative effect of children to 3.4%, although the effect remains statistically significant. In Model 2, each of the job interruptions measures is statistically significant, suggesting that relative to women who do not experience any job interruptions, those who do have interruptions experience a penalty to their wages that increases with the number of weeks of absence per year, and this effect holds across types of interruption. Thus, both the job interruptions measures *and* the number of children produce negative wage effects for women.

In Model 3, human capital measures are included (highest grade completed, tenure at current job, and labor force experience), which reduces the magnitude of the negative effect of children on women's wages. The effects of the job interruption measures remain significant. Though the effects are slightly reduced, the pattern remains consistent: relative to persons who do not have absences, those who have absences incur wage penalties that increase with the number of absences. In Model 4, measures of the respondent's current job are included, and again, further reduce, but do not eliminate the negative relationship between children and women's wages. Controlling for human capital, job interruptions, and job characteristics, women receive a 1.5% penalty to their wages for every child in their household. Negative effects of job interruptions also remain.

For men, results reported in Table 3 indicate a marriage premium of between 5% and 7%, while there is no significant effect of children in any of the models tested. Further, it appears that men get a "pass" for job interruptions due to family reasons. In Model 2, the relationship between men's job interruptions and wages is similar to that of women's: Compared to men who do not have any job absences, those who do incur wage penalties that increase with increased incidents of absence for any type of interruption.

Table 3 about here

In Model 3, when human capital measures are included, the negative effect of 12 or fewer weeks of job interruptions due to family reasons is reduced to non-significance, while the magnitude of the effects of other job interruption are reduced, but not eliminated. Thus, Model 3 suggests that men who have 12 or fewer weeks of absence from work due to family reasons experience wage outcomes no different from men who have no such absences. However, men who have more than 12 weeks of family-related job interruptions experience a 3.5% decline in their wages relative to men who have no such absences.

In Model 4, when measures of the current job are included, the negative effects of job interruptions due to family reasons are completely eliminated, while the other types of job interruptions remain negatively associated with wages. This final model suggests that net of human capital and job characteristics, men receive a "pass" for job interruptions that occur for family reasons, but are penalized for other types of job interruptions. Moreover, men receive greater penalties to their wages for non-family and involuntary job absences than women receive for the same types of job absence. However, women also receive a penalty to their wages for job absences due to family reasons, net of the penalty to their wages associated with the presence of children.

Long-term Effects of Leave-taking: Estimating Wage Trajectories.

Next, fixed-effects estimates are applied to the partial adjustment model. The model provides estimates of an ultimate expected "equilibrium" wage, as well as the relative speed at which that wage is achieved. The equilibrium wage represents the predicted wage at which a given individual is expected to plateau as a function of specified parameters. Using the fixed effects estimates of model parameters on the rate of wage change is calculated. This “dynamic coefficient” represents the speed of achievement of the equilibrium wage, and is a function of the affect of wages in one year on wages in the next. While this estimate cannot be meaningfully interpreted on its own, it does provide a point of comparison between men’s and women’s wage process. The greater the absolute value of the dynamic coefficient, the greater the speed with which wages achieve equilibrium.

Employees who have job absences may derail themselves from a wage-earning path. Workers may realize higher returns to job experience had a leave not been taken from their job. Similarly, workers who have no job absences may be better positioned for promotions or other types of rewards that are realized over time. Job interruptions may depress an employee's earnings trajectory relative to employees who did not take leave and were not so derailed. Results reported above suggest that this does occur in the short-term. The next section examines effects of interruptions over the working life course. Employer responses to job absences may be subtle and cumulative over time. Detecting such outcomes requires longitudinal data over several years. The method of analysis to be employed here allows the estimation of a rate of change in wages over time, which captures the dynamics of job histories over the working life course, and enables the estimation of a wage trajectory.

RESULTS: Women’s and Men’s Wage Trajectories

Table 4 contains the dynamic coefficients and substantive parameters describing wage trajectories over time. The fixed effect estimates from which the substantive parameters in Table 4 are derived appear in Appendix A. The estimates presented here differ from the fixed effects estimates represented in Tables 2 and 3, as these are derived from an equation that includes a lagged Y term.

Table 4 about here

The dynamic coefficient contained in the lagged value of the dependent variable indicates that women reach their equilibrium wage value at a faster rate than men do[4].

Figure 3 shows predicted wages for men and women up through the fifteenth year of labor force participation[5]. According to the predicted wage growth pattern, by approximately the fifth year of work, both women and men have nearly reached their plateau in wages, at which point, wage growth is negligible. Figure 3 depicts a lower equilibrium wage for women relative to men, and also illustrates sex differences in the process of wage growth, such that men not only enter the labor market at higher initial wages, but the gap in earnings between men and women widens over the working life course. And although men reach their equilibrium wage at a slower pace than women do[6], men may ultimately expect to earn higher hourly wages than women may ($13.08 for men versus $12.05 for women).

The "long-run" parameters generated by the partial adjustment model (Table 4) suggest that among men, having up to 12 weeks of voluntary or involuntary job absence is associated with a decrease of approximately 2% per hour in their equilibrium wage, relative to the peak wages of men who have no such absence. Having more than 12 weeks of voluntary absence is associated with a decrease of 3% in equilibrium wages. Twelve or more weeks of involuntary absence exacts an hourly penalty of 5% to the equilibrium wage for men. Over the long run, married men receive a wage premium of approximately $1.05 per hour[7].

Among women, absence from work for family reasons is not associated with any long-term wage penalties. Like men, women also receive a 2% decrease in wages if up to 12 weeks of work are missed for voluntary, non-family, or involuntary reasons. However, women who have more than 12 weeks of voluntary or involuntary absences experience wage outcomes no different from women who have no such absences. The presence of each child exacts a wage penalty of approximately $0.98 per hour on women's equilibrium wage, while marriage is associated with a $1.00 increase in equilibrium wages.

Discussion

Although women experience immediate wage penalties associated with all types of job interruptions, over the long run, only non-family types of job absences up to 12 weeks a year affect wages. Men's wages on the other hand, are immediately affected by all non-family job interruptions, and these effects do persist over the long run, depressing men's equilibrium wages as well. Moreover, the wage gap between men who have job interruptions and men who do not is larger than the gap between women who have job absences and women who do not, and this hold for both short- and long-term wage consequences. This suggests that men's absence from work is less acceptable than is women's.

The marriage premium for men is confirmed in these findings, and men receive this premium in the short- and long-run. Women receive a marriage premium as well, but the effect is considerably smaller. The presence of children remains a penalty to women's wages both in the short- and long-run, while men receive no such penalty. That this motherhood penalty persists net of job absences, tenure, labor force experience, and other human capital characteristics points to two possibilities: 1) motherhood is a stigmatized status in the workplace and women are penalized for their motherhood status, or 2) motherhood is exhausting and women burdened with a "second shift" reduce their productivity at work either to reserve energy for the second shift, or as a result of burnout from the second shift. The former explanation would seem more plausible if wage penalties associated with family job absences were found. However, since job absences associated with family reasons do not affect women's wages over the long run, employer bias seems less plausible.

That men receive a wage premium for marital status, however, does speak to potential bias in the workplace that favors married men and discriminates against women with children. Alternatively, the wage premium could reflect the privileges married men receive from the housework and care provided by a wife. In addition, the marriage premium may also reflect increased productivity among married men who may have greater incentive to work hard and strive for raises and/or promotions to financially support their families.

Men and Family Leave

My initial suspicion with these results was that the leaves that men are reporting are actually associated with family reasons for the absence, but are only being reported as a non-family leave for fear of employer or other social reprisal for transgressing a normative expectation. Studies find that men who take leave following the birth of a child generally do not avail themselves of formal parental leave policies. Rather, they employ a mix of vacation days, sick days, and "other" or "personal" days (Hyde, et al. 1993; Pleck 1993). Men may be reluctant to take family leave, and/or reluctant to label a leave as such when one is actually taken. Instead, men may label their absence a "sick leave" or "vacation", even though the time taken off corresponds with the birth of a child.

Thus, I suspected that non-family leaves reported by men in the NLSY(a break from work due to respondent illness or inability to work, the respondent was on vacation, did not feel like working, or “other”) may actually have been leaves taken to care for family needs surrounding the birth of a child. However, investigation into this possibility did not support my contention. Although there was some evidence of misreporting, accounting for the misreports did not change the findings in any way.

Conclusions

Although men may indeed be reluctant to take family leave, results presented here suggest there are no wage penalties specifically associated with family leaves from employment. These findings suggest that men and women should feel free to take advantage of family leave policy provisions. For women, while there are immediate wage penalties associated with taking non-family leave, these effects do not persist over the long run.

While my expectation that men would experience greater wage penalties associated specifically with family leave-taking was not supported, I do find evidence that over the long run, the process of wage growth differs for men and women and wage penalties are distributed differently; women’s wages are more sensitive to the presence of children, whereas men’s wages are more sensitive to job absence. Further, women appear to be rewarded for tenure on the job, but are not negatively affected by taking leave, while men are not rewarded for tenure on the job, and are penalized for involuntary and non-family leaves. Thus, women are rewarded for staying with an employer, while men are penalized for absence from employment. Men may find greater success in moving through successively more lucrative jobs, and so are not rewarded for tenure, while women may find greater rewards through staying at the same job.

That the wage penalties associated with men’s job leaves are not found in family leave-taking suggests that men actually get a “pass” for absences from work due to family reasons. No similar distinction operates for women’s leave-taking. Penalties associated with women’s wages do not seem to be about job discontinuity, or even about “signaling” associated with job leaves. Rather, penalties to women’s wages come directly from the presence of children.

Thus, this finding provides further evidence of the “baby tax”, “motherhood penalty”, or “family gap” that has been consistently reported in the recent literature. The effect is net of leaves taken, so the human capital argument that the motherhood penalty comes from an inconsistent work history is not supported. Rather, results suggest that women with children suffer a wage penalty either due to employer discrimination or to decreased productivity at work as a result of having to work the “second shift” at home.

Since I find no effect of job leaves on women’s wages, it does not appear that employers associate leaves with negative evaluations of women as workers that would result in wage penalties. Given that employers do not penalize women for taking leave, it seems illogical to expect that they would discriminate against women for their parental status if there are no penalties associated with interruptions from work associated with parental status. Thus, the more likely explanation for the child penalty to women’s wages is that women are overburdened with having to juggle demands at work and demands at home because the primary care for children remains the sole province of women. Time-use data would help to clarify this conjecture. Unfortunately, the NLSY data do not provide such information.

As long as the penalty for men’s leave-taking, whether real or imagined, persists, women will likely remain disproportionately burdened with the wage penalties associated with family responsibilities, and the gender gap in wages will continue. Until men become willing and able to participate more fully in family responsibilities such as primary childcare, and relieve their wives of some of the physical, psychic, and emotional work inhered in such labor, the motherhood penalty will likely persist. Results reported here suggest that men should take advantage of family leave policies to do just that, without fear of negative repercussions on their wages. Work poses obstacles to men’s family involvement, and to overlook these is to rest the issue of work-family conflict squarely and unjustly on women’s shoulders, thereby perpetuating gender inequality both in the labor market as well as in the family.

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Figure 1: Predicted Hypothetical Wage Trajectories for Men and Women.

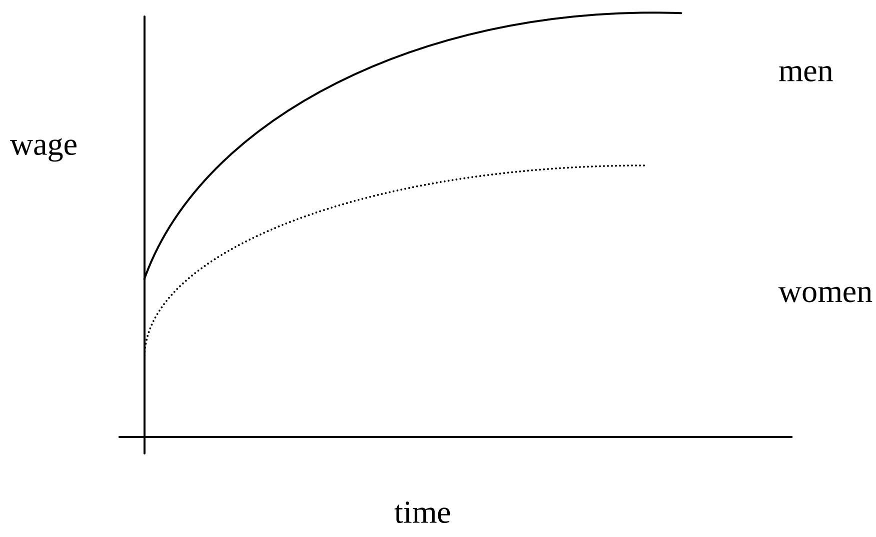

Figure 2. Trajectory of Y Over Time.

Derived from dY/dt = bY + aX when initial level of Y (Y(0) is below equilibrium (Y*).

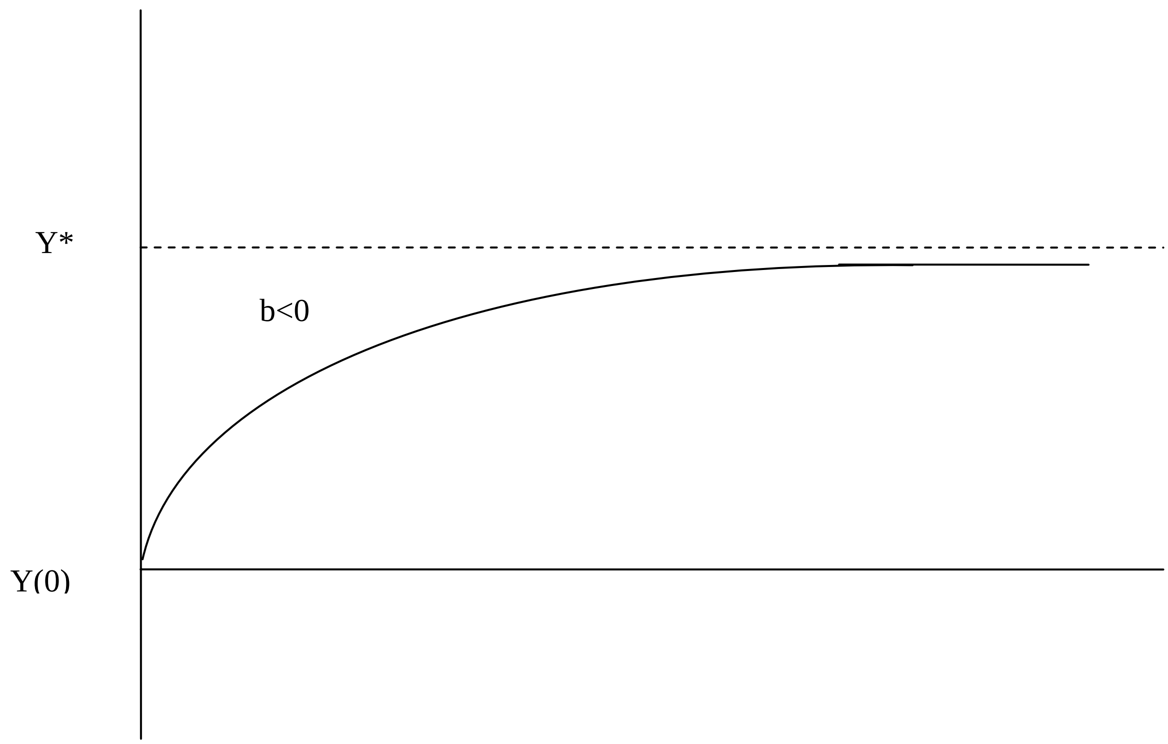

Figure 3. Predicted Wage Profiles by Sex

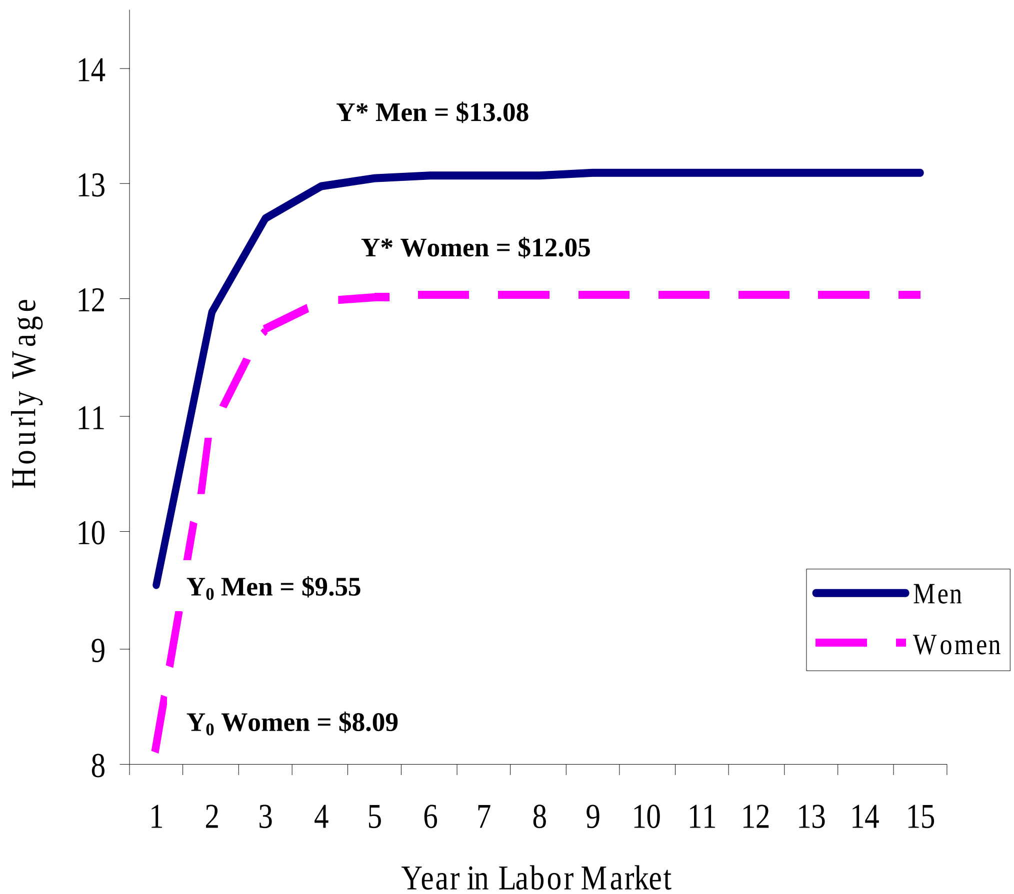

Table 1. Means and Standard Deviations of Variables in Models by Sex: NLSY 1979-1998

Variables	**Men (N=4,618)**		**Women (N=4,586)**		**Total (N=9,204)**	
	First Year	Last Year	First Year	Last Year	First Year	Last Year
Hourly wages	9.23 (1.72)	13.45 (1.90)	7.69 (1.66)	10.65 (1.89)	8.41 (1.70)	11.97 (1.92)
Weeks of family job interruptions	0.82 (4.33)	0.57 (3.88)	5.54 (12.25)	3.17 (9.73)	3.17 (9.48)	1.86 (7.51)
Weeks of voluntary non-family job interruptions	8.07 (12.57)	1.29 (5.83)	10.13 (14.22)	1.82 (7.06)	9.10 (13.46)	1.55 (6.48)
Weeks of involuntary job interruptions	10.16 (14.14)	3.58 (9.93)	8.46 (12.75)	2.85 (8.71)	9.31 (13.48)	3.22 (9.35)
Age	21.87 (2.77)	35.09 (3.95)	22.37 (3.13)	35.07 (4.20)	22.12 (2.96)	35.08 (4.07)
Number of children	0.24 (0.60)	1.50 (1.37)	0.57 (0.97)	1.64 (1.31)	0.40 (0.83)	1.55 (1.34)
Marital Status (Proportion Married)	0.19 (0.40)	0.70 (0.46)	0.38 (0.48)	0.79 (0.41)	0.28 (0.45)	0.74 (0.44)
Highest grade completed	12.34 (2.40)	12.86 (2.52)	12.61 (2.20)	13.20 (2.39)	12.48 (2.31)	13.03 (2.46)
Proportion in Professional Occupations	0.16 (0.36)	0.30 (0.46)	0.17 (0.38)	0.33 (0.47)	0.17 (0.37)	0.32 (0.47)

Table 1 continued

Variables	**Men (N=4,574)**		**Women (N=4,542)**		**Total (N=9,115)**	
	First Year	Last Year	First Year	Last Year	First Year	Last Year
Weeks Tenure	69.45 (65.05)	282.27 (265.13)	64.00 (58.00)	254.36 (248.79)	66.74 (61.69)	268.34 (257.47)
Months of Full Time Labor Force Experience	74.41 (67.73)	533.25 (234.63)	61.17 (60.57)	411.76 (244.41)	67.80 (64.59)	472.96 (247.10)
Months of Part Time Labor Force Experience	10.29 (33.18)	46.20 (76.84)	17.65 (41.14)	104.22 (133.41)	13.96 (37.55)	75.01 (112.48)
Duncan SEI Score	29.48 (21.28)	37.09 (23.79)	38.38 (20.91)	44.74 (22.10)	33.92 (21.56)	40.90 (23.28
Proportion in Union	0.20 (0.40)	0.05 (0.22)	0.15 (0.35)	0.05 (0.22)	0.17 (0.38)	0.05 (0.22)
Proportion Working Full Time	0.85 (0.35)	0.91 (0.29)	0.75 (0.43)	0.74 (0.44)	0.80 (0.40)	0.83 (0.38)
Disability	0.04 (0.19)	0.04 (0.19)	0.06 (0.24)	0.06 (0.23)	0.05 (0.21)	0.05 (0.21)

Table 2. Fixed Effects Regression Models of Log of Hourly Wages by Sex: NLSY 1979-1998: Women *p<.05, **p<.01, ***p< .001, two-tailed. Unstandardized OLS coefficients. Standard errors in parentheses.

	Model 1	Model 2	Model 3	Model 4
Variables at t-1				
Age	.026*** (.001)	.024*** (.001)	.003 (.002)	.003 (.002)
Marital Status	.028*** (.007)	.030*** (.007)	.026*** (.007)	.024*** (.007)
Number of Children	-.038*** (.004)	-.034*** (.004)	-.017*** (.004)	-.015*** (.004)
Job Interruptions[a]:				
Family –12 weeks or less		-.017** (.006)	-.016*** (.006)	-.015* (.006)
Family – greater than 12 weeks		-.045***(.008)	-.039***(.008)	-.032***(.008)
Voluntary, Non-Family 12 weeks or less		-.029*** (.006)	-.024*** (.006)	-.023*** (.006)
Voluntary, Non-Family greater than 12 weeks		-.039***(.007)	-.031***(.007)	-.027***(.007)
Involuntary – 12 weeks or less		-.030*** (.006)	-.022*** (.006)	-.020** (.006)
Involuntary – greater than 12 weeks		-.049***(.007)	-.036***(.008)	-.031***(.008)
Highest Grade Completed			.048*** (.003)	.045*** (.004)
Weeks Tenure at Current Job			1.01E-04***(.001)	9.35E-05***(.001)
Months of Full Time Labor Force Experience			4.68E-04*** (.001)	4.69E-04*** (.001)
Months of Part Time Labor Force Experience			1.50E-04*** (.001)	1.64E-04** (.001)
Professional Occupation				.030*** (.007)
Duncan SEI Scores				.001*** (.001)
Full Time Employment				.018** (.006)
Union Status				.054*** (.006)
Disability				-.026** (.009)
Intercept	1.60*** (.013)	1.67*** (.013)	1.46*** (.052)	1.43*** (.053)
R^2	.068	.085	.229	.252

a: The reference group for each of the job interruptions dummy variables is “no interruptions”.

Table 3. Fixed Effects Regression Models of Log of Hourly Wages by Sex: NLSY 1979-1998: Men *p<.05, **p<.01, ***p< .001, two-tailed. Unstandardized OLS coefficients. Standard errors in parentheses.

	Model 1	Model 2	Model 3	Model 4
Variables at t-1				
Age	.021*** (.001)	.019*** (.001)	-.007** (.002)	-.007** (.002)
Marital Status	.072*** (.006)	.065*** (.006)	.060*** (.006)	.054*** (.006)
Number of Children	-.005 (.003)	-.004 (.004)	-.005 (.003)	-.005 (.003)
Job Interruptions[a]:				
Family –12 weeks or less		-.019* (.009)	-.014 (.006)	-.014 (.009)
Family – greater than 12 weeks		-.045** (.016)	-.035* (.016)	-.026 (.016)
Voluntary, Non-Family 12 weeks or less		-.029*** (.006)	-.027*** (.006)	-.026*** (.006)
Voluntary, Non-Family greater than 12 weeks		-.068*** (.008)	-.059*** (.008)	-.051*** (.008)
Involuntary – 12 weeks or less		-.037*** (.005)	-.030*** (.005)	-.028*** (.005)
Involuntary – greater than 12 weeks		-.072*** (.006)	-.065*** (.006)	-.060*** (.006)
Highest Grade Completed			.058*** (.004)	.053*** (.004)
Weeks Tenured at Current Job			6.77E-05*** (.001)	6.19E-05*** (.001)
Months of Full Time Labor Force Experience			5.70E-04*** (.001)	5.96E-04*** (.001)
Months of Part Time Labor Force Experience			4.14E-04*** (.001)	4.69E-04** (.001)
Professional Occupation				.003 (.008)
Duncan SEI Scores				.001*** (.001)
Full Time Employment				.037*** (.007)
Union Status				.078*** (.006)
Disability				-.029* (.010)
Intercept	1.88*** (.012)	1.96*** (.013)	1.75*** (.057)	1.73*** (.057)
R^2	.073	.091	.219	.238

a: The reference group for each of the job interruptions dummy variables is “no interruptions”.

Table 4. Estimated 'Long-Run' Parameters for Partial Adjustment Model of Change. Log of Hourly Wages by Sex: NLSY 1979-1998.

Variables at t-1	Men	Women
Hourly Wage Dynamic Coefficient	-1.20***	-1.38***
<u>Long-run</u> <u>Effects of Independent Variables:</u>		
Age	-.004	.007**
Married	.051***	.018*
Number of children	-.009	-.013*
1-12 Weeks of Family Job Interruptions	-.011	-.015
12+ Weeks of Family Job Interruptions	-.025	-.021
1-12 Weeks of Voluntary, Non-Family Job Interruptions	-.023**	-.023**
12+ Weeks of Voluntary, Non-Family Job Interruptions	-.033**	-.013
1-12 Weeks of Involuntary Job Interruptions	-.020**	-.020**
12+ Weeks of Involuntary Job Interruptions	-.053***	-.021
Highest grade completed	.056***	.046***
Weeks tenured	-1.96E-05	2.87E-05
Months of Full Time Labor Force Experience	5.52E-04***	4.09E-04***
Months of Part Time Labor Force Experience	3.72E-04**	8.58E-05***
Professional Occupation	-.003	.024**
Duncan SEI scores	5.13E-04*	6.01E-04**
Full Time Employment	.079***	.034***
Union Status	.053***	.043**
Disability	-.030*	-.031*
Intercept	1.63***	1.36***

Results of significance tests from Appendix A. *p<.05 **p<.01 ***p<.001, two-tailed.

Appendix A. Fixed Effects Regression Estimates of Log of Hourly Wages by Sex Including Lagged Values of Dependent Variable: NLSY 1979-1998.

Variables at t-1	Men	Women
Natural Log of Hourly Wage	.301*** (.005)	.252*** (.005)
Age	-.003 (.002)	.005** (.002)
Married	.036*** (.006)	.013* (.007)
Number of children	-.006 (.003)	-.009* (.004)
1-12 Weeks of Family Job Interruptions	-.008 (.008)	-.011 (.006)
12+ Weeks of Family Job Interruptions	-.017 (.016)	-.016 (.008)
1-12 Weeks of Voluntary, Non-Family Job Interruptions	-.016** (.006)	-.017** (.006)
12+ Weeks of Voluntary, Non-Family Job Interruptions	-.023** (.007)	-.009 (.007)
1-12 Weeks of Involuntary Job Interruptions	-.014** (.005)	-.015** (.006)
12+ Weeks of Involuntary Job Interruptions	-.037*** (.006)	-.016 (.007)
Highest grade completed	.039*** (.003)	.034*** (.003)
Weeks tenured	-7.66E-06 (.001)	2.10E-05 (.001)
Months of Full Time Labor Force Experience	3.86E-04*** (.001)	3.06E-04*** (.001)
Months of Part Time Labor Force Experience	2.60E-04** (.001)	6.42E-05*** (.001)
Professional Occupation	-.002 (.007)	.018** (.007)
Duncan SEI scores	3.59E-04* (.000)	4.50E-04** (.001)
Full Time Employment	.055*** (.007)	.026*** (.005)
Union Status	.037*** (.005)	.032** (.006)
Disability	-.021* (.010)	-.023* (.009)
Intercept	1.14*** (.056)	1.02*** (.052)
R^2	.514	.475

*p<.05, **p<.01, ***p< .001, two-tailed. Standard errors in parentheses.

Endnotes

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[1] Women in the United States earn approximately 84 cents on every dollar earned by men (U.S. Department of Labor 2002).

[2] Budig and England (2001) used the 1982-1993 waves of the NLSY. Newmark and Korenman (1994) used the NLS Young Women 1973-1982 waves and Lundberg and Rose (2000) used the Panel Study of Income Dynamics (PSID) 1980-1992. Each of these studies employs a pooled cross-section time-series method in their analyses.

[3] Y* represents the equilibrium value of hourly wage.

[4] Dynamic coefficient = |ln(lagged dependent variable)|.

[5] Predicted wages are calculated with the following equation derived from the difference equation:

$$Y_t = e^{bt}\, Y_0 + (1 - e^{bt})\, (Y^*)$$

where t represents years in the labor market, Y* is the equilibrium wage, b is the dynamic coefficient, and Y_0 is the starting wage. The initial starting wage is generated from a linear regression equation of wages at the end of the second year on covariates in the first year of work.

[6] The absolute value of men's dynamic coefficient |ln(b*)|= 1.20, while women's dynamic coefficient = 1.38, indicating men's slower relative speed to equilibrium.

[7] Wage penalties are calculated by exponentiating the long run parameters from the partial adjustment model. For example, the long run coefficient for marital status among men is .051. Exp(.051) = 1.05.